THE COST OF JOB LOSS AND THE "NEW" PHILLIPS CURVE

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INTRODUCTION

Almost two decades after Lucas and Sargent's [1981] famous, but premature, requiem for the short-run Phillips curve, Fuhrer could claim that it was still "alive and well, and living in a good number of … macroeconometric models" [1995, 41]. However much the theoretical foundations, or for that matter the practical consequences, of this "inexorable and mysterious tradeoff" [Mankiw, 2000] had been debated in the interim, it survived as an "extremely robust *empirical* relationship, [one] showing little or no sign of instability over … 35 years" [Furher 1995, 41, emphasis added].[1] In the brief time since then, however, public and professional confidence in the Phillips curve, even as a limited empirical phenomenon, has waned. Given the recent combination of low, and more or less stable, rates of wage and price inflation in the United States, and a remarkable decline in the unemployment rate, to its lowest level in more than three decades, this was perhaps inevitable. There is anecdotal evidence, for example, that even the Federal Reserve has "unlearn[ed] this lesson of Econ 101" [Nasar 1998]: Alan Greenspan is sometimes identified as a "new paradigm" advocate, and the Open Market Committee has often seemed more concerned with "irrational exuberance" in financial markets and measures of *product* market slack.[2]

The compromise position [Gordon, 1998; Blanchard, 2000] is that problems with the Phillips curve are limited to price-based specifications: nominal wages, it is said, did rise faster over the second half of the last decade than the first—3.3 percent per year from 1991 to 1995, versus 4.1 percent between 1996 and 2000—even as the annual rate of price inflation fell, from 3.3 to 3.1 percent. An "expectations augmented" version of Phillips' [1958] original wage-based specification has even experienced a theoretical rehabilitation of sorts as the adjustment mechanism [Roberts, 1995; Blanchard and Katz, 1999] consistent with new Keynesian "wage curves."[3] It is our position, however, that the differences in the wage- and price-based Phillips curves have been overstated. In particular, while no one doubts that the "inflation shortfall" [Gordon, 1998] is more substantial for prices than wages, out-of-sample forecasts of

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nominal wage inflation for the last decade predict more labor market "pressure" than has been observed. It is this shortcoming that our paper is intended to address.

The two most common explanations for the seemingly anomalous behavior of wages and prices are NAIRU drift[4] and/or the establishment of a new paradigm.[5] There are notable exceptions, of course: the so-called "Goldilocks interpretation" [Gordon, 1998] attributes much of the inflation shortfall to a sequence of fortuitous aggregate supply shocks and the mismeasurement of prices.[6] Furthermore, of these two, economists have devoted (far) more attention to the former. Katz and Krueger [1999], for example, have underscored the effects of an older, and therefore less mobile, labor force, increased reliance on temporary workers, and the rise in the incarceration rate on the NAIRU and, as an indirect consequence, nominal wage behavior. In a similar vein, Gordon [1997] himself had earlier identified "labor peace," a decrease in the real minimum wage and the decline in labor's share in national income as three potential sources of NAIRU variation.

A reasonable list of other recent "labor market anomalies" with potential consequences for desired real wages and/or nominal wage adjustments would include the reduction in job security for some classes of workers [Jaegar and Huff Stevens, 1999], the "changing face of job loss" [Farber, 1997], and workers' increased *fears* of displacement, well-founded or otherwise [Aaronson and Sullivan, 1998]. Of particular relevance for this paper, such a list would also include the *persistence* of jobless spells. Valletta [1998] and Needels and Nicholson [1999], for example, have observed that mean duration did not fall as much in the most recent expansion as past experience would have once led one to predict. It has often proven difficult, however, to isolate the effects of these on labor market outcomes: as Wanner characterizes the current literature, "where there was once anecdotal smoke, there is now some statistical fire, but no raging conflagration" [1999, iii].

The effects of "globalization" on American wages have also drawn some attention. There is little disagreement, for example, that free(r) trade flows have reduced, at least in relative terms, the demand for low-skill workers in the United States. The size of this effect, however, is controversial [Wood, 1995; Freeman, 1995], and the implications of this and other "sectoral shifts" for aggregate labor market outcomes remain uncertain.[7]

Our own research is intended to complement this now diverse literature, and is premised on the existence of a third, and perhaps much simpler, alternative to NAIRU drift and new paradigms. In particular, we explore the notion that *if* a stable wage-based Phillips curve does exist, it is better modelled using another "metric" of labor market slack, one whose behavior better reflects recent labor market developments. In other words, we concur with Brainard's observation that the "original motivation … [for the Phillips curve was] … that unemployment was a good measure of pressure in labor markets" [1998, 340], but observe that the phrase "a good measure" does not preclude the existence of other, flow-based, measures, or the construction of Phillips-inspired models based on them.[8] To be more specific, we find that substitution of a normalized "cost of job loss" measure for the unemployment rate in Blanchard and Katz's [1999] model of nominal wage inflation produces comparable full sample fits and better out-of-sample forecasts for the recent past. In other words, we find that standard models cannot "explain" the recent past without recourse to régime shifts, and that our alternative can.

In the next section, we introduce the alternative measure of labor market slack, and provide some intuition for our empirical results. The third section outlines a common framework for the comparison of the two measures—the Blanchard and Katz [1999] model. In the two subsequent sections, we discuss full sample estimates and their implications, the results of several diagnostic checks, and out-of-sample forecast properties. We draw four conclusions from the data. First, consistent with Gordon [1998] and Blanchard [2000], the evidence in favor of a Phillips-consistent relationship between nominal wage inflation and labor market slack, however defined, is stronger than the conventional wisdom holds, even without the now common additions of an "error correction term" and the growth of output per worker. Second, absent substantial and implausible NAIRU drift, traditional models predict nominal wage inflation in excess of that observed over the last decade, which is consistent with the presence of misspecification problems. Third, the alternative model produces better out-of-sample forecasts of wage inflation for the same period, which hints that the misspecification involves omitted, or incorrect, variables, rather than NAIRU drift or other kinds of régime shifts. And fourth, it is difficult to rationalize the existence of a stable, or even predictable, NAIRU within either model, a result with important implications for the conduct of stabilization policies.

THE UNEMPLOYMENT RATE AND THE COST OF JOB LOSS

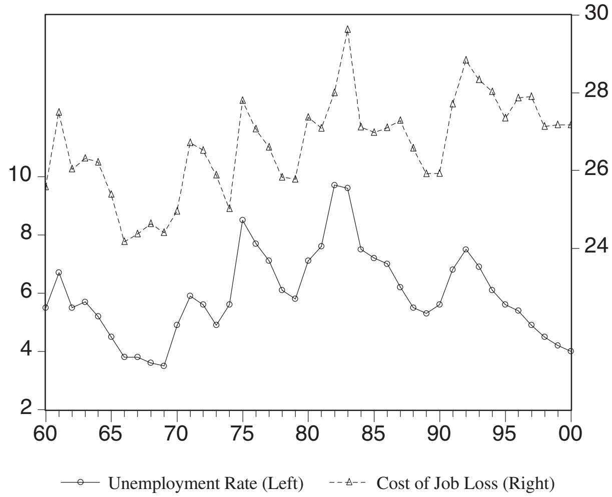
Schor and Bowles [1987] define the one-period real cost of job loss C_t to be:

(1) $$C_t = \omega_t - (q_t\omega_t^U + (1-q_t)\omega_t^R)$$

where ω_t, ω_t^U and ω_t^R are, respectively, the annual wages of the representative job holder, the value of unemployment insurance and other income replacement benefits and the re-employment wage, all measured in constant dollars, and where q_t is the expected duration of unemployment for job losers, *expressed as a fraction of one year*.[9] Inasmuch as the value of C_t determines the "default position" in bargaining models where labor market transactions are "contested" [Bowles and Gintis, 1993]—a class that includes both mainstream and heterodox specifications[10]—it will influence the behavior of firms and workers more than the *number* of those out of work. Workers who expect to search longer for new offers, or who suspect that offers to "outsiders" will become less attractive in the future, will moderate their wage demands, even if the unemployment rate has not changed. More important in the present context, falling unemployment need not generate wage and price pressures if the cost of job loss—in particular, the duration of jobless spells—does not follow suit.

For our purposes, we found it useful to calculate the cost of job loss on a per week basis, and to express the result as a fraction of average earnings per week in the covered sector. To provide readers with a sense of the relevant magnitudes, the mean value of our "normalized cost of job loss," c_t, over the full sample period was 26.7 percent; that is, the average cost of job loss just exceeded one quarter of "pre-separa-

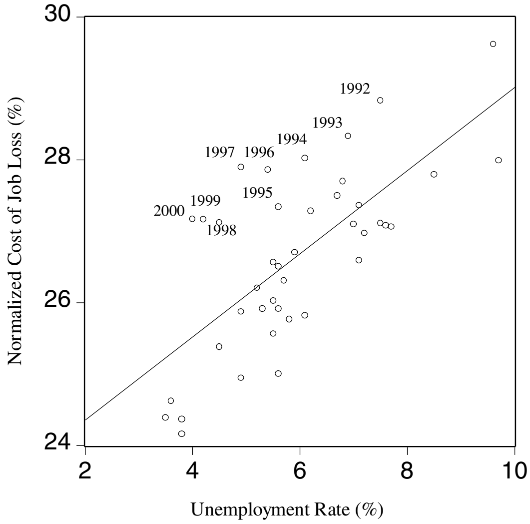
FIGURE 1
Unemployment and the Normalized Cost of Job Loss in the United States, 1960-2000



tion" earnings. These calculations involve (at least) two simplifications, however. Given limited aggregate time series data on the re-employment wage, we followed the lead of Schor and Bowles [1987], who assume on the basis of cross-sectional studies that $\omega_t^R = 0.87\,\omega_t$.[11] Since there is little reason to suppose that the "incumbent's premium," the percentage difference between ω_t^R and ω_t, has increased over the last decade (increased turbulence in labor markets [Aaronson and Sullivan, 1998] would lead one to suspect the opposite, in fact) we are confident that this simplification is not responsible for the observed persistence of the cost of job loss. Data constraints were also responsible for a second simplification: unlike Schor and Bowles [1987], we do not attempt to construct a separate duration series q_t for job losers, but substitute the mean duration of unemployment insurance collection. Further details and sources are included in an appendix, and the full data set is available on request from the authors.

With this in mind, Figure 1 is a simple time series plot of the normalized cost of job loss c_t and the unemployment rate u_t over the full (1960-2000) sample period. The data exhibit two important properties. First, until the 1991-92 recession, the two series move more or less in unison, *modulo* a small phase shift: c_t follows u_t with a small, almost imperceptible, lag. For as long as this pattern prevailed, the two measures were in some sense equivalent, and specification of the Phillips curve was, at least in this sense, moot. Second, this equivalence has since broken down, so that researchers must now (re)consider how best to measure labor market pressure. Over the last decade, for example, the unemployment rate fell, from 6.8 percent to 4.0 percent, but the cost of job loss remained more or less constant, at close to 27 percent. There are smaller intervals in which the cost of job loss even rose: between 1990 and 1995, for example, u_t neither rose nor fell (5.6 percent) but c_t increased from 25.9

FIGURE 2
Umemployment and the Normalized Cost of Job Loss in the United States, 1960-2000



percent to 27.3 percent. The broader picture, however, is one in which the recent and persistent decline in the rate of joblessness has had little effect on the cost of job loss.

To provide another perspective on the behavior of these series, Figure 2 is a scatter plot of the same data, together with a simple (bivariate OLS) best-fit line. Despite the considerable "tug" of recent observations on the line, *every observation since the last recession is an outlier of sorts* in the sense that the cost of job loss was higher, and sometimes much higher, than was once consistent with recent unemployment rates. The recent behavior of c_t hints, therefore, that labor markets are "looser" than sometimes claimed.

Further examination of the raw data leads to the conclusion that the culprit is the surprising persistence of jobless spells, which remained close to their pre-expansion levels until the late 1990s. Figure 3 is a scatter plot of mean unemployment insurance collection in weeks and the unemployment rate over the same period, with the corresponding best-fit line. Once more, almost all recent observations are outliers, in which duration is "too long." Between 1996 and 2000, for example, mean duration was 14.3 weeks, about the same as it was between 1976 and 1980 (14.1 weeks) even though the unemployment was more than two percentage points lower (4.6 percent versus 6.7 percent). As late as 1996, duration was close to 15 weeks, a level experienced just four times in our sample (1975, 1976, 1980 and 1983) before the 1990s. The bottom line is that even in the current "high pressure labor market" [Katz and Krueger, 1999], job losers must still wait a substantial time for a reasonable offer. The intuition for our empirical results is that this has served to moderate wage demands.

We recognize, of course, that no one measure of labor market slack is ideal: the persistence of mean duration, for example, does not tell us much about its *distribution*. It is possible, for example, that spells for all but a small, but increasing, number

FIGURE 3
Unemployment Rates and the Duration of Jobless Spells in the United States, 1960-2000

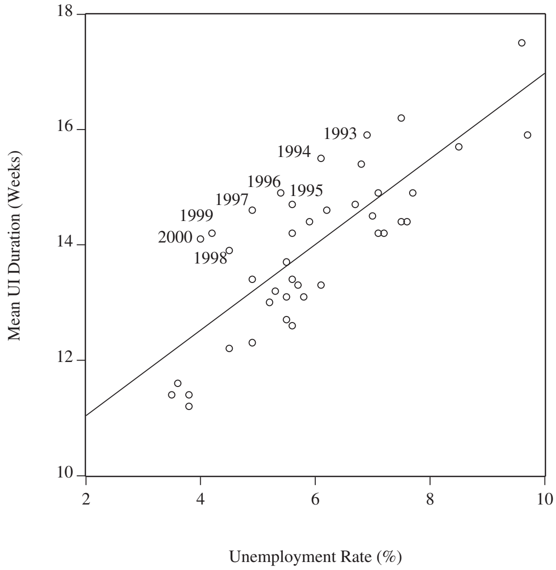
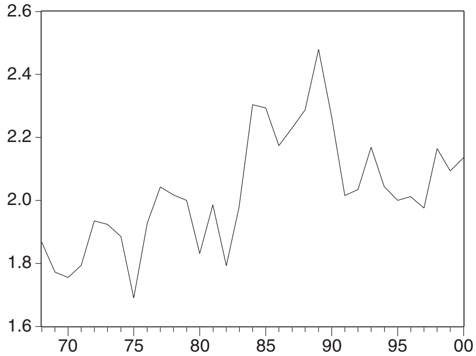


FIGURE 4
The Ratio of the Mean to Median Jobless Spell in the United States, 1968-2000



of "hard core jobless" have decreased, in which case our measure would understate "true" labor market pressure. To this end, consider the behavior of the ratio of the mean to median jobless spell, plotted in Figure 4 for the shorter period over which median data is available. While this ratio is larger now than it was during the 1970s, it is also smaller than it was during the 1980s, from which we conclude that distributional shifts are not responsible for the divergence. This said, the mean spell has risen, relative to the median, since 1998, consistent with a decline in the relative bargaining power of unskilled workers.

A FRAMEWORK FOR COMPARISON

The Blanchard and Katz [1999] model, one of several recent attempts to build sturdier microfoundations for the Phillips curve, provides a convenient framework for the comparison of the two measures of labor market slack. "Under some … assumptions about functional form *and the appropriate indicator of labor market tightness*" [Blanchard and Katz 1999, 70, emphasis added], the core of their model is a simple wage curve:

$$w_t - p^e_t = \mu b_t + (1-\mu) y_t - \beta u_t + e_t \tag{2}$$

where w_t and p^e_t are the logarithms of the nominal wage and expected price level in period t, b_t and y_t are the logarithms of the mean reservation wage and output per worker in the same period, u_t is the unemployment rate and e_t is an error term that could well be autocorrelated. In behavioral terms, workers demand a nominal wage such that their expected real wage is a weighted average of the reservation wage and labor productivity, adjusted in some loose sense for variations in bargaining power, as embodied in the unemployment rate. As Blanchard and Katz [1999] note, this specification encompasses a number of recent models, including, for example, Shapiro and Stiglitz [1984] ($\mu = 1$) and Mortensen and Pissarides [1994] ($\mu = 0$).

The existence of a Phillips-based relationship follows from the assumed sluggishness of the reservation wage:

$$b_t = a + \lambda(w_{t-1} - p_{t-1}) + (1 - \lambda) y_t \tag{3}$$

The rationale for equation (3) is that the reservation wage depends, in the first instance, on workers' aspirations and the real value of unemployment insurance benefits, both of which are sensitive to real wages in the (immediate) past. It will also be a function of output per worker to the extent that the value of the alternatives available to workers (leisure, home production, the informal sector, *et cetera*) and nonlabor income are.

Substitution of equation (3) into equation (2) and some manipulation of terms then produces:

$$\Delta w_t = a\mu + (p^e_t - p_{t-1}) - (1 - \lambda\mu)(w_{t-1} - p_{t-1} - y_{t-1}) + (1 - \lambda\mu)\Delta y_t - \beta u_t + e_t \tag{4a}$$

where the third term, the difference between last period's real wage and output per worker, is understood to be an "error correction term," a now common specification [OECD, 1997]. To estimate equation (4a), the inflation rate in the previous period, $p_{t-1} - p_{t-2}$, is substituted for the otherwise unobserved expected inflation rate for the current period, $p^e_t - p_{t-1}$, with two important caveats. First, we do *not* claim that these adaptive expectations are exploitable. That is, since we do not characterize the empirical Phillips curve in terms of a menu of choices, a menu that would no doubt be affected if treated as such, the force of the Lucas critique is somewhat blunted. Second, and perhaps more important, the substitution of actual inflation forecasts, based on the Federal Reserve Bank of Philadelphia's Livingston survey, does not alter our basic conclusions, as detailed in the next section.

If, as Blanchard and Katz [1999] then assert, $\mu\lambda \cong 1$ for the United States—but not Europe—equation (4a) becomes the familiar textbook Phillips curve [Blanchard, 2000; Dornbusch, Fischer and Startz, 2000]:

$$\Delta w_t = a\mu + (p^e_t - p_{t-1}) - \beta u_t + e_t. \tag{4b}$$

The structure of the alternative model is almost identical, and is based on the modified wage curve:

$$w_t - p^e_t = \mu^* b^*_t + (1-\mu^*)y_t - \beta^* c_t + e^*_t \tag{5}$$

where c_t is the normalized cost of job loss, as defined in the previous section, and b^*_t is that part of the reservation wage not included in the calculation of c_t. The distinction between b^*_t and b_t is needed because the direct effects of variations in unemployment insurance benefits, for example, are included in the costs of dismissal, while the direct effects of variations in extra-market opportunities are not. The determination of the modified b^*_t series then mimics equation (3):

$$b^*_t = a^* + \lambda^*(w_{t-1} - p_{t-1}) + (1-\lambda^*)y_t \tag{6}$$

so that the alternative Phillips curve is:

$$\Delta w_t = a^*\mu^* + (p^e_t - p_{t-1}) - (1-\lambda^*\mu^*)(w_{t-1} - p_{t-1} - y_{t-1}) + (1 - \lambda^*\mu^*)\Delta y_t - \beta^* c_t + e^*_t. \tag{7a}$$

The "textbook variant" of equation (7a) is then:

$$\Delta w_t = a^*\mu^* + (p^e_t - p_{t-1}) - \beta^* c_t + e^*_t . \tag{7b}$$

Our comparison of c_t and u_t is then based on a comparison of equations (4a) and (4b) with equations (7a) and (7b).

It should be noted that within this framework both the standard and alternative models are *linear* in their respective measures of labor market slack. This is in contrast with traditional or textbook representations in which the short-run Phillips curve is drawn convex to the origin or, in behavioral terms, in which inflation is assumed to accelerate as unemployment falls, and *vice versa*. Some [Akerlof *et al.*, 1996] have even concluded that *both* the short- and long-run Phillips curves are convex in low inflation economies. On the other hand, Eisner [1997, 1998] concludes on the basis of local labor market behavior that it could be concave, or even S-shaped. Gordon tests for both sorts of curvature, and concludes that the "short run Phillips curve is resolutely linear, at least within the range of inflation and unemployment observed over the [sample] period" [1997, 28]. Even if Gordon [1997] is not correct, we believe that our results do not turn on the use of a linear approximation, but leave a more complete discussion of this issue for future research.

FULL SAMPLE ESTIMATION

We estimated six versions of the two models with four decades of annual[12] data from 1960 to 2000: an unrestricted general specification, in which neither of the restrictions[13] embodied in equations (4a) and (7a) were imposed; a restricted general specification, in which both restrictions were imposed, two partially restricted general specifications, in which one restriction, but not the other, was imposed; an unrestricted textbook specification, in which no restriction is imposed on the inflation coefficient in equations (4b) or (7b); and a restricted textbook specification, in which the coefficient is set equal to one. We calculated Δw_t as the difference in the logarithms of the average weekly wage in covered employment; $p^e_t - p_{t-1}$, as the lagged difference in the logarithms of the new CPI-U series; $z_t = w_{t-1} - p_{t-1} - y_{t-1}$, as the lag of the logarithm of the share of wages and salaries in national income; and Δy_t, as the annual rate of growth of the Bureau of Labor Statistics output per hour index for nonfarm business, all multiplied by 100.[14] Consistent with Fuhrer [1995] and others, we expected to find, and correct for, limited first-order autocorrelation, but in most cases, these corrections had little effect on the parameter estimates. These estimates were obtained using a nonlinear least-squares algorithm, rather than the Cochrane-Orcutt or similar procedure, and are asymptotically equivalent to maximum likelihood estimates. Tables 1 and 2 summarizes our results, including the F-statistics and relevant p-values for each set of restrictions.

It is the similarities, not the differences, in the full-sample properties of the estimated models that are most remarkable, consistent with the close movement of c_t and u_t until the 1990s. First and foremost, in both models, the unrestricted general specification, one of the partially restricted general specifications, and the unrestricted textbook specification "explain" substantial, and almost identical, proportions of the variation in nominal wage inflation over the last four decades. The adjusted R^2 for the unrestricted standard model, for example, is 0.68, and that for the unrestricted alternative is 0.67. Likewise, both unrestricted textbook models have an even higher adjusted R^2 of 0.70.

Second, there is little evidence that both restrictions are sensible in either model, and less so in the alternative: in all cases, the *p*-value is less than one in a thousand. The data also reveal that the problem is the restriction on lagged inflation $p_{t-1} - p_{t-2}$, the *p*-values for which are similar. The superficial explanation is that $p_{t-1} - p_{t-2}$ is a poor substitute for $p^e_t - p_{t-1}$, perhaps because, to use an old fashioned term, expectations about inflation are inelastic, a feature that could well vanish if exploited. Under this interpretation, a one percentage-point increase in the rate of price inflation in the current period causes expectations for the next period to be revised upward about six tenths of one percentage point in both models and across most specifications. To further explore this issue, one could either estimate a separate model of price infla-

TABLE 1
Estimates for the Standard Phillips Curve, 1960-2000

$$\Delta w_t = \pi_1 + \pi_2(p_{t-1} - p_{t-2}) + \pi_3 z_t + \pi_4 \Delta y_t + \pi_5 u_t + e_t$$
$$e_t = \rho e_{t-1} + u_t$$

	Restrictions					
	None	$\pi_2 = 1$	$\pi_3 + \pi_4 = 0$	$\pi_2 = 1$, $\pi_3 + \pi_4 = 0$	$\pi_3 = \pi_4 = 0$	$\pi_2 = 1$, $\pi_3 = \pi_4 = 0$
π_1	6.71[b]	2.38	6.74[b]	0.80	4.67[a]	5.70[a]
π_2	0.62[a]	(1.00)	0.60[a]	(1.00)	0.63[a]	(1.00)
π_3	0.04	−0.06	0.04	−0.10		
π_4	0.02	0.24[c]	−0.04	0.10		
π_5	−0.41[b]	−0.86[a]	−0.40[b]	−0.88[a]	−0.46[b]	−0.88[a]
ρ	0.11	0.22	0.10	0.31[c]	0.13	0.32[b]
adj R^2	0.68	0.54	0.69	0.53	0.70	0.54
SE	1.02	1.22	1.01	1.23	0.99	1.23
SSR	35.2	52.1	35.6	54.7	36.1	57.2
log L	−54.2	−62.1	−54.4	−63.0	−55.5	−65.0
DW	1.93	1.90	1.94	1.89	1.93	1.87
F statistic		16.3	0.30	9.67	0.38	6.83
p-value		0.00	0.59	0.00	0.69	0.00

The columns correspond to the unrestricted general model, the (two) partially restricted general models, the restricted general model, the unrestricted textbook model, and the restricted textbook model.
a. Significant at the 1 percent level; b. significant at the 5 percent level; c. significant at the 10 percent level.

TABLE 2
Estimates for the Alternative Phillips Curve, 1960-2000

$$\Delta w_t = \pi_1 + \pi_2(p_{t-1} - p_{t-2}) + \pi_3 z_t + \pi_4 \Delta y_t + \pi_5 u_t + e_t$$
$$e_t = \rho e_{t-1} + u_t$$

	Restrictions					
	None	$\pi_2=1$	$\pi_3 + \pi_4=0$	$\pi_2 = 1$, $\pi_3 + \pi_4=0$	$\pi_3 = \pi_4 = 0$	$\pi_2 = 1$, $\pi_3 = \pi_4 = 0$
π_1	13.7[b]	13.7[b]	13.3[b]	9.28	13.3[a]	15.3[b]
π_2	0.55[a]	(1.00)	0.60[a]	(1.00)	0.54[a]	(1.00)
π_3	−0.01	−0.19[c]	0.04	−0.23[b]		
π_4	0.09	0.44[b]	−0.04	0.23[b]		
π_5	−0.45[b]	−0.88[a]	−0.40[b]	−0.77[a]	−0.41[c]	−0.55[b]
ρ	0.09	0.27	0.10	0.41[b]	0.05	0.48[b]
adj R^2	0.67	0.39	0.69	0.38	0.70	0.35
SE	1.03	1.41	1.01	1.43	1.00	1.46
SSR	36.1	69.4	35.6	73.1	37.1	81.3
log L	−54.7	−67.8	−54.4	−68.8	−56.2	−72.2
DW	1.89	1.81	1.94	1.80	1.85	1.79
F statistic		31.4	0.39	18.7	0.21	15.8
p-value		0.00	0.53	0.00	0.81	0.00

The columns correspond to the unrestricted general model, the (two) partially restricted general models, the restricted general model, the unrestricted textbook model, and the restricted textbook model.
a. significant at the 1 percent level; b. significant at the 5 percent level; c. significant at the 10 percent level.

TABLE 3
Standard and Alternative Phillips Curve Estimates with Livingston Forecast Series, 1960-2000

$$\Delta w_t = \pi_1 + \pi_2(p^e_t - p_{t-1}) + \pi_3 z_t + \pi_4 \Delta y_t + \pi_5 u_t\ (\text{or } \pi_5 c_t) + e_t$$

$$e_t = \rho e_{t-1} + u_t$$

			Restrictions			
	None	None	$\pi_3 + \pi_4 = 0$	$\pi_3 + p_4 = 0$	$\pi_3 + p_4 = 0$	$\pi_3 + p_4 = 0$
π_1	10.8[a]	16.0[a]	11.3[a]	15.1[a]	3.88[a]	12.3[a]
π_2	0.74[a]	0.63[a]	0.63[a]	0.55[a]	0.64[a]	0.57[a]
π_3	0.15[a]	0.09	0.14[b]	0.10		
π_4	0.13	0.15	−0.14[b]	−0.10		
π_5 (u_t)	−0.43[a]		−0.33[b]		−0.33[c]	
π_5 (c_t)		−0.38[b]		−0.29[c]		−0.38[b]
ρ	−0.03	0.07	−0.02	0.05	0.30[c]	0.17
adj R^2	0.73	0.70	0.70	0.68	0.67	0.67
SE	0.94	0.99	1.00	1.03	1.04	1.04
SSR	30.1	33.5	34.7	37.1	39.1	38.9
log L	−51.1	−53.2	−53.9	−55.2	−56.3	−56.2
DW	1.99	1.90	1.96	1.92	2.01	1.91

a. Significant at the 1 percent; b. significant at the 5 percent; c. significant at the 10 percent.

tion and substitute its *ex ante* forecasts for the expectations series, or use *bona fide* expectations data.[15] We chose the latter but found that when $p^e_t - p_{t-1}$ was set equal to the mean inflation forecast of those in the Livingston survey group, the results were not much affected, as summarized in Table 3. In particular, most estimated coefficient values, including those on price inflation, were similar across specifications, with little or no improvement in fit.

Our results were also robust with respect to the choice of a consumer price index. If, for example, the CPI-U-X1 is substituted for the CPI-U in the general unrestricted versions of either model, there are no substantive differences in either estimated parameters or significance values.[16]

Third, without the problematic restriction on lagged inflation, and consistent with Blanchard and Katz's [1999] characterization of the empirical literature, the coefficients on the error correction term and the growth of output per hour often had the wrong signs but were insignificant. To be more precise, we find support in both models for the null hypothesis that the sum of the coefficients is zero and for the (joint) null hypothesis that each is zero or, in terms of the notation of the previous section, $\mu\lambda = 1$. The latter implies, and the estimates of the final specifications confirm, that the textbook models are reasonable approximations, at least for the United States.

Fourth, in all cases, the estimated coefficient on labor market pressure, measured in terms of either c_t or u_t, is significant in both the economic and statistical senses, but sensitive to the (still suspect) restriction on lagged inflation. Without this restriction, a one percentage-point increase in the unemployment rate is estimated to reduce the rate of wage inflation about four tenths of a percentage point in the standard model, with or without the second restriction. The estimated effect of a one percentage-point increase in the normalized cost of job loss is almost identical, which does not come as much of a surprise: when c_t is regressed on u_t over the full sample period, the estimated slope coefficient, allowing for first-order serial correlation, is 0.87. Neglecting their eventual divergence, then, a one percentage-point increase in the unemployment rate is associated with an almost one percentage point increase in the normalized cost of job loss, so that a one percentage point increase in either should have a similar effect on nominal wage inflation.

Last, if there is a stable NAIRU, neither model is robust enough to produce reliable, or even sensible, estimates of it. As Blanchard and Katz [1999] demonstrate for the standard model, if one defines x_t, a measure of market power, to be the difference between the logarithms of output per worker y_t and the expected real wage $w_t - p_t^e$, and then assumes that (i) x_t and y_t are constant and (ii) expectations are "correct," $p_t^e = p_t$, equations (2) and (3) are then consistent with a NAIRU u_N whose value is:

$$u_N = \beta^{-1}[\mu a + (1-\lambda\mu)x]. \tag{8a}$$

Even within this austere framework, variations in "market conditions" will cause the NAIRU to drift unless $\mu\lambda = 1$. On the other hand, if, as our empirical work suggests, $\mu\lambda = 1$ is reasonable, the values of x_t and y_t do not matter—and therefore need not be assumed constant—and the estimated NAIRU assumes the familiar form of the ratio of the constant to the (remaining) slope coefficient:

$$u_N = \beta^{-1}\mu a. \tag{8b}$$

The implied NAIRU point estimates for the two textbook specifications are 10.1 percent in the unrestricted case, and 6.5 percent in the restricted case, both of which are too large, even as full sample estimates. One could likewise define a *NAICJL*, the value of which would be:

$$c_N = \beta^{*-1}[\mu^* a^* + (1-\lambda^*\mu^*)\, x]. \tag{9a}$$

or, in the textbook case $\mu\lambda = 1$,

$$c_N = \beta^{*-1}\mu^* a^* \tag{9b}$$

Our point estimates of c_N are 32.4 percent in the unrestricted textbook specification, and 27.8 percent in the restricted specification, which are no more plausible than the estimates of u_N.

NAIRU advocates will perhaps counter that our failure to build one or more breaks into our models has contaminated these estimates. The surprise, perhaps, is that it is more difficult to document, and then date, such breaks than conventional wisdom holds. To see this, consider the cumulative sum of standardized residuals or CUSUM [Brown, Durbin and Evans, 1975] for the unrestricted general specifications of both models, plotted in Figures 5 and 6, together with the relevant 5 percent significance bands. If the parameter estimates are stable from period to period, the CUSUM should not wander far from zero—that is, remain between the two significance bands. As

FIGURE 5
Cumulative Sum of Standardized Residuals (CUSUM) and 5 Percent Significance Bands for the Unrestricted Standard Model

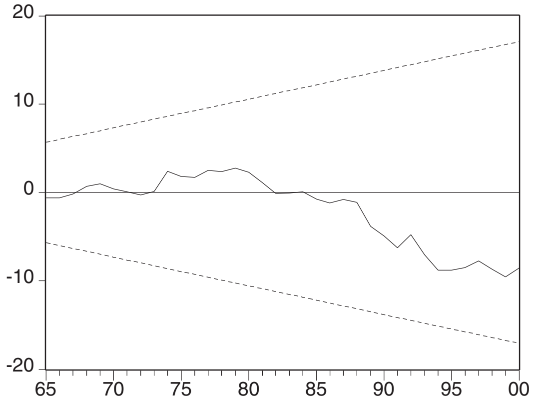
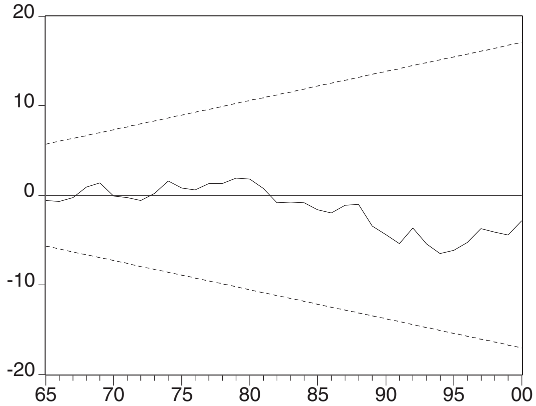


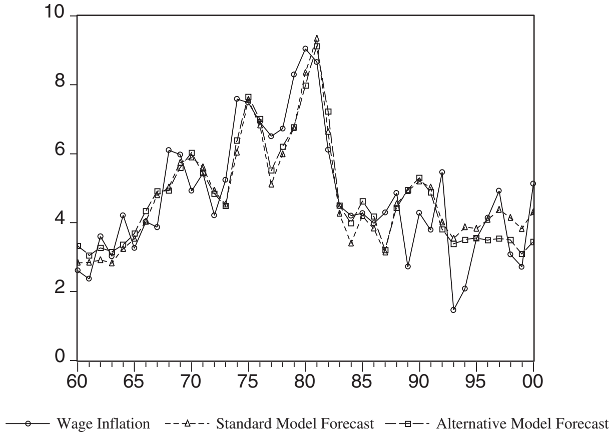
FIGURE 6
Cumulative Sum of Standardized Residuals (CUSUM) and 5 Percent Significance Bands for the Unrestricted Alternative Model



Harvey [1990] and others have noted, however, CUSUM plots are most useful as an informal diagnostic tool: sharp fluctuations constitute *prima facie* evidence of structural breaks, while persistent drift signals an omitted variable(s) or other misspecification problem. It is therefore important to note that in both models, the CUSUMs remain well within the 5 percent bands over the entire sample and, furthermore, exhibit no telltale spikes.

This said, the CUSUM for the standard model *does* start to drift downward sometime in the late 1980s, while the CUSUM for the alternative model remains more or less constant. This hints at the existence of a choice, between a modified standard model, one that must treat the last decade as anomalous, and an alternative model

FIGURE 7
Ex Post Dynamic Forecasts of Wage Inflation Using the Standard and Alternative Textbook Models



that does not need to be modified. We shall consider this choice in more detail in the next section, but note that Occam's Razor would favor the latter.

FORECASTING: DO WE NEED A NEW TEXTBOOK MODEL?

The emergence of a choice, between a standard model that needs to be modified, and an alternative that does not, is underscored when we compare forecast performance, both *ex ante* and *ex post*. Consistent with the parametric tests in the previous section, we focus on the "textbook" versions of both models. Consider first the *ex post* dynamic forecasts of the unrestricted (coefficient on lagged price inflation) specifications, as depicted in Figure 7. Neither model captures all of the variation in nominal wage inflation over the first third of the sample, and each tends to overstate the amount of variation over the middle third. Both also miss the slowdown in wage inflation in the late 1980s. The forecasts differ most, however, from 1992 onward. The standard model captures some of the acceleration in wages between 1995 and 1997, but almost none of the decelerations between 1993 and 1994, and 1998 and 1999. As a result, the mean forecast for the period from 1993 to 2000, 3.99 percent, is almost 20 percent higher than the observed 3.38 percent. The alternative model, on the other hand, captures less of the variation but the mean forecast, 3.43, is just 1.5 percent too high.

A comparison of the *ex ante* forecasts for the same period leads to similar conclusions. We re-estimated the restricted and unrestricted textbook specifications of both models over the sub-sample from 1960 to 1992, and constructed dynamic out-of-sample forecasts for the period from 1993 to the present. Figures 8 and 9 plot these forecasts, and Table 4 summarizes their properties—root mean square error (RMSE), mean absolute error (MAE), mean absolute percentage error (MAPE), and the Theil statistic, as well as the decomposition of the forecast error into bias, variance and covariance proportions.[17]

FIGURE 8
Ex Ante Forecasts of the Standard Textbook Model(s), 1993-2000

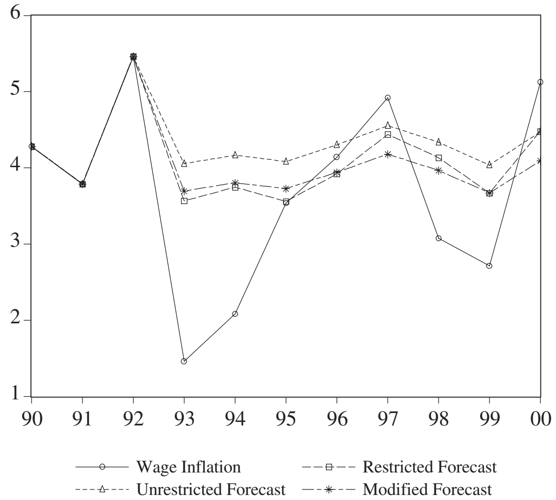


FIGURE 9
Ex Ante Forecasts of the Alternative Textbook Model(s), 1993-2000

Armed with either standard model and knowledge of the subsequent behavior of price inflation and unemployment at the end of 1992, a macroeconomist would have predicted that nominal wage inflation would hover around 4 percent for the remainder of the decade, a rate well in excess of that observed on all but two occasions, 1997 and 2000. With the unrestricted alternative textbook model, on the other hand, the same macroeconomist would have failed to predict the subsequent variation in nominal wage inflation, but also would have been correct *on average*. In particular, she would have overpredicted the amount of labor market pressure in 1993, 1994, 1998, and 1999, underpredicted it in 1996, 1997 and 2000, and been on the mark in 1995, with positive and negative errors of about the same size. Furthermore, if she relied on the restricted version, she would have been surprised at how *much* wage inflation

TABLE 4
Forecasting Nominal Wage Inflation with Standard, Alternative and Modified Standard Textbook Phillips Curves, 1993-2000

	Standard Unrestricted	Standard Restricted	Alternative Unrestricted	Alternative Restricted	Modified Standard
RMSE	1.38	1.11	1.26	1.75	1.19
MAE	1.12	0.89	1.05	1.58	1.00
MAPE	50.9	40.3	42.8	51.3	43.2
Theil's U	0.18	0.15	0.17	0.29	0.16
Decomposition					
Bias	0.40	0.25	0.03	0.27	0.18
Variance	0.56	0.61	0.72	0.17	0.77
Covariance	0.04	0.14	0.25	0.56	0.06

there was, despite the low, and falling, rate of unemployment: in this case, the forecasts fall short from 1995 onward.

The measures of forecast performance in Table 4 provide some more perspective on these observations. Both versions of the standard textbook model and the unrestricted alternative textbook model have similar RMSE, MAE and MAPE—the restricted standard model's are a little bit lower than the unrestricted alternative model, and the unrestricted standard model's are a little bit higher—but there are dramatic differences in the composition of the forecast error. In heuristic terms, the proportion due to bias, which reflects differences in the means of the actual and forecast series and is most indicative of "fundamental" shortcomings, should be as small as possible. Conditional on the bias proportion, the proportion due to variance, which reflects differences in variation of the two series, should also be small, too. On the other hand, the proportion due to covariance, which reflects "non-fundamental" error, should be the largest of the three. It is therefore important to note that in both the restricted and unrestricted versions of the standard textbook model, the proportions due to bias and variance "explain" almost all the forecast error. In the unrestricted, for example, the bias and variance proportions are 40 percent and 56 percent, respectively. In contrast, in the unrestricted alternative model, the proportion due to bias is less than 5 percent, and while the proportion due to variance (72 percent) exceeds that in either version of the standard textbook model, so does the covariance proportion (25 percent). Despite their similar RMSEs, therefore, it seems clear that the unrestricted alternative textbook model is "better" than either version of the standard model.

Indeed, it is not even clear that despite its inferior RMSE—.75, which is 0.37 or 20 percent more than the next highest— the restricted alternative textbook model is the poorest of the four. The bias proportion (27 percent) is about the same as that in the restricted standard textbook specification (25 percent) and much smaller than in the unrestricted (40 percent), and the variance proportion (17 percent) is the smallest of the four. It is in fact the one specification in which most of the forecast error is not "fundamental."

Last, we note that the unrestricted alternative model has much in common with a "modified standard" model. Suppose, for the sake of argument, that in 1992, the same macroeconomist has noticed the CUSUM drift beginning in 1987, and introduces a shift term into the unrestricted standard textbook specification, consistent the reservation wage $b_t = a_0 + a_1 d_{87} + \lambda(w_{t-1} - p_{t-1}) + (1 - \lambda)y_t$, where d_{87} assumes the value 0 before 1987, and 1 after. The forecast series is (also) plotted in Figure 8, and its behavior seems much closer to the unrestricted alternative model, an impression consistent with the performance data in Table 4. Its RMSE, MAE and MAPE are similar to those of the unmodified standard model, but the bias proportion is smaller, if still higher than that in the unrestricted alternative. This is sufficient, in our view, to establish a rough equivalence between the modified standard and unmodified alternative models and, to the extent that the latter is the simpler of the two, to favor the latter.

CONCLUSION

It will perhaps surprise some readers that despite these results, we remain agnostic about the existence of a stable Phillips curve of either the traditional or alternative varieties. We *do* believe, however, that some, perhaps most, modern theories of wage determination are consistent with the use of the cost of job loss as a sensible measure of labor market pressure. This observation has important implications for the specification and estimation of macroeconomic models, one of which we have considered in this paper: in a model that features wage curves and the adjustment of nominal wages to these, substitution of this alternative measure enhances our understanding of recent labor market behavior. We do not believe, however, that the cost of job loss captures all that is "anomalous" about the current environment, and expect that subsequent research will lead to better, perhaps composite, indices.

APPENDIX

The raw data used in the paper were:

W_t: average weekly wage in covered employment, from the Department of Labor's Employment and Training Administration Handbook No. 394 and recent UI Data Summaries, both available at *workforcesecurity.doleta.gov/unemploy/datastats.asp*.

P_t: CPI-U for all urban consumers, from the Bureau of Labor Statistics (BLS) web site, *stats.bls.gov*.

P^a_t: alternative price level, CPI-U-X1, from Table B-62, Economic Report of the President 2001, *w3.access.gpo.gov/eop*.

P^f_t: mean 12-month December price level forecast of Livingston group, based on P^o_t, both available from the FRB Philadelphia web site, *www.phil.frb.org/econ/liv*.

d^{UI}_t: mean actual duration of UI claims in weeks, from the Department of Labor's Employment and Training Administration Handbook No. 394 and recent UI Data Summaries, both available at *workforcesecurity.doleta.gov/unemploy/datastats.asp*.

d^{MN}_t: mean duration of unemployment for workers 16 and older in weeks, from the Bureau of Labor Statistics (BLS) web site, *stats.bls.gov*.

d^{MD}_t: median duration of unemployment for workers 16 and older in weeks, from from the Bureau of Labor Statistics (BLS) web site, *stats.bls.gov*.

B_t: average weekly UI benefit amount, from the Department of Labor's Employment and Training Administration Handbook No. 394 and recent UI Data Summaries, both available at *workforcesecurity.doleta.gov/unemploy/datastats.asp*.

F_t: total family assistance, in billions, from Table B-29, Economic Report of the President 2001, *w3.access.gpo.gov/eop*.

N_t: total civilian employment, in billions, Table B-36. Economic Report of the President 2001, *w3.access.gpo.gov/eop*.

u_t: unemployment rate, from the BLS web site, *stats.bls.gov*.

Y_t: index of output per hour in non-farm business, from Table B-49, Economic Report of the President 2001, *w3.access.gpo.gov/eop*.

S_t: total wage and salary disbursements, from Table B-27, Economic Report of the President 2001, *w3.access.gpo.gov/eop*.

I_t: national income, in billions, from Table B-27, Economic Report of the President 2001, *w3.access.gpo.gov/eop*.

The constructed series were:

$$w_t - w_{t-1} = 100[\log(W_t) - \log(W_{t-1})]$$

$p_{t-1} - p_{t-2}$ = either $100[\log(P_{t-1}) - \log(P_{t-2})]$ or $100[\log(P^a_{t-1}) - \log(P^a_{t-2})]$, as context requires

$$p^e_t - p_{t-1} = 100[(P^f_t - P^o_t)/P^o_t]$$

$$y_t - y_{t-1} = 100[(Y_t - Y_{t-1})/Y_{t-1}]$$

$$q_t = d^{UI}_t/52$$

$$A_t = F_t/52N_t$$

$W^u_t = B_t + A_t$, average weekly value of UI and other income-replacement benefits

$CJL_t = W_t - [q_t W^u_t + (1-q_t)0.87\, W_t]$, nominal cost of job loss, measured on a per week basis

$$c_t = CJL_t/W_t, \text{ normalized cost of job loss}$$

$$z_t = 100^*\log(S_t/I_t)$$

NOTES

We thank Middlebury College for generous leave support (Matthews), the Departments of Economics at Yale and UC San Diego (Matthews), where this paper was revised, and the Christian A. Johnson summer internship program (Kandilov). We have also benefited from conversations with David Colander, Carolyn Craven, Paul Sommers and Paul Wonnacott, and from the detailed comments of two reviewers. The usual caveat holds.

1. This verdict was never unanimous, of course. Our colleagues Sommers and Malinov [1997], for example, discern evidence of a stable price-based Phillips curve for most OECD countries *except* the United States. It was, however, commonplace, at least until the current "new era." There are important differences between our respective papers, however, not least of which is Sommers and Malinov's [1997] focus on price, as opposed to wage, inflation. To the extent that the relationship between these has become more complicated and/or less stable, comparisons are difficult to draw. For more, see Gordon [1998] and Blanchard [2000].
2. See, for example, Greenspan's remarks at the CEPR at Stanford in 1997. The full text is available at *www.federalreserve.gov/boarddocs/speeches/19970905.htm*.
3. While the term "wage curve" is owed to Blanchflower and Oswald [1994], heterodox economists since at least Desai [1975] and Rowthorn [1977] have devoted considerable attention to the relationship between real wages and various measures of bargaining power, including the rate of unemployment, and the adjustment of nominal wages to such "proto-wage curves." For reviews of the broader literature, see, for example, Carlin and Soskice [1990] or Jossa and Musella [1998].
4. We have adopted Tobin's [1997] distinction between the NAIRU and the so-called "natural" rate, and we focus on the former.
5. It is unfortunate, in our view, that recent discussions of either Phillips curve are often tied to still unresolved debates [Galbraith, 1997; Stiglitz, 1997] over the existence, and possible variation over time, of the natural rate or NAIRU. Even in the otherwise mainstream models estimated here, for example, the connection between Phillips-consistent nominal wage behavior and the natural rate is more flexible than sometimes supposed.
6. Gordon [1998] identifies a mixture of both "traditional" and "new" AS shocks as important, where the former includes food, energy and import prices and the latter the prices of computer services and medical care. The now familiar claim [Boskin et al., 1998] that fluctuations in the standard CPI overstated the "true" rate of price inflation implies that the shortfall in price inflation is smaller than first seems, but the implications for wage inflation are less clear. In our models, CPI bias matters because it affects the calculation of real wages and therefore our estimates of wage curves. Following current BLS practice, we deflate nominal wages with the chained CPI-U, but as a robustness check, we also report results for the CPI-U-X1. This said, we are confident, based on the work of Abraham et al. [1998] and others, that the bias is smaller than the Boskin Commission concluded.
7. In still unpublished work, Ewing and Wunnava [2000] reject the presence of a unit root in North American jobless rates in favor of a trend that "breaks" with the imposition of NAFTA.
8. To illustrate the role of labor market flows in wage bargaining, consider two economies in which 5 million workers in a labor force of 100 million are without work each period. In the first, however, no more than a few thousand are "separated" from firms each period and an equal number are hired, but in the second, the number separated and hired is close to a million. It follows that the principal stock measure of labor market slack, the rate of unemployment, will be constant, and equal, in the two economies, but that most flow measures—the likelihood of rehire in a particular period, for example—will not be. It is our contention that the two economies will not have the same "wage setting curve" [Layard et al., 1991].
9. It follows that q_t is related to, but not the same as, the more familiar likelihood of rehire. This one-period measure is a theoretical and empirical convenience, but one that Schor and Bowles [1987] found reasonable. If the real wages of (re)hired workers return to pre-separation levels one period later, of course, the one-period and multi-period measures are identical: when q_t is 0.25, for example, the worker who loses her current position can expect to receive ω_t^U for three months, ω_t^R for nine months and ω_t thereafter, in which case the cost of job loss is the difference between $0.25\ \omega_t^U + 0.75\ \omega_t^R$ and ω_t in the current period, and 0 in each successive period.

10. The responses of Williamson [1993] and Stiglitz [1993] to the Bowles and Gintis [1993] paper, for example, suggest that this approach is consistent with both "transactions costs" and the "new economics of information."
11. For some additional evidence that the assumed proportion (0.87) is a sensible one, see also Chowdhury and Nickell [1985] and Blanchflower [1991].
12. Our rationale for the use of annual data is that it remains common practice, and allows for more immediate comparisons with "textbook representations" and/or conventional wisdom. Given the recent debate(s) between Sommers and Malinov [1999, 2000] and Ewing and Seyfried [1999] and Payne, Tracy and Potter [2000], we were also concerned that the use of more frequent observations could obscure as much as it illuminated.
13. The first of these restrictions is that the sum of the coefficients on the error correction term $w_{t-1} - p_{t-1} - y_{t-1}$ and the growth of output per worker per week Δy_t is zero, and the second is that the coefficient on lagged inflation $p_{t-1} - p_{t-2}$ is one.
14. The construction of the c_t and u_t series were discussed in the second section. Measurement concerns prompted the substitution of a direct measure of labor's share.
15. There is another alternative—substitution of actual for expected inflation on the premise that expectations are "rational"—but this introduces other problems, both econometric and methodological.
16. The CPI-U-X1 is based on the rental equivalence method for homeowner costs for consumer prices before 1983, when this method was first adopted. When this substitution is made, the estimated unrestricted regression models were:

$$\begin{array}{l} \Delta w_t = \underset{(2.3)}{7.02} + \underset{(7.0)}{0.77}\Delta p_{t-1} + \underset{(.84)}{0.05}z_t + \underset{(0.47)}{0.06}\Delta y_t - \underset{(-3.2)}{0.51}u_t \\ \text{adj } R^2 = 0.70 \qquad \log L \ \nabla -53.3 \qquad DW = 1.93 \end{array}$$

and

$$\begin{array}{l} \Delta w_t = \underset{(3.5)}{14.2} + \underset{(6.7)}{0.64}\Delta p_{t-1} - \underset{(-.04)}{0.002}z_t + \underset{(0.75)}{0.11}\Delta y_t - \underset{(-2.5)}{0.47}c_t \\ \text{adj } R^2 = 0.57 \qquad \log L = -55.0 \qquad DW = 1.87 \end{array}$$

with t-statistics in parentheses. A third "research series," the CPI-U-RS, is now available for the period from 1980 onward, but not before then. For details, see the most recent Economic Report of the President, available online at *w3.access.gpo.gov/eop*.
17. For a review of forecast evaluation, including the shortcomings of this decomposition of the forecast error, see Theil [1966] and Kennedy [1992].

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